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ABSTRACT

The impact of the timing, spacing, and number of children on a married woman's wage growth over her life cycle was examined. The data used for the analysis were information pertaining to the labor market experience of women and the birth dates of their children, taken from the 1976 survey of the Michigan Panel Study of Income Dynamics (IDP). There were four principal findings. First, the number of children a woman rears has an important deterrent effect on her accumulated work experience. A woman who has one child will average about 2.5 years less work experience over her lifetime than a woman with no children. A second child results in an additional two years withdrawn from the labor market. Second, larger birth intervals (holding constant the number and timing of children) are associated with larger reductions in lifetime work effort. The magnitude of this effect appears to increase with the number of children. Third, the timing of the first birth (the age at which a woman has her first child) does not appreciably affect lifetime labor market experience. Fourth, child rearing, through its effect on accumulated work experience, has a substantial effect on wage rates. (EM)

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FAMILY FORMATION, LABOR MARKET EXPERIENCE, AND THE WAGES OF MARRIED WOMEN

PREPARED FOR THE NATIONAL INSTITUTE OF CHILD HEALTH AND HUMAN DEVELOPMENT, NATIONAL INSTITUTES OF HEALTH

JOHN F. COGAN, FRANKLIN BERGER

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PREFACE

The research presented in this report was funded by the National Institute of Child Health and Human Development under Contract No. N01-HD-52837-NIH. This contract funded research into the impact of the spacing, timing, and number of children in a family on several aspects of family behavior. Among those examined are:

- The effects of parity, number of siblings, and uncertainty on the *educational attainment of children*, in which both numbers of children and investments in children are influenced by parental choice.
- The effects of children on parent economic well-being over time. This work considers both *consumption and savings patterns* over the life cycle and the relationship between a *wife's earnings* and her prior (and expected future) fertility experience.

This report examines the impact of the number, timing, and spacing of births on the mother's life-cycle patterns of labor market work and wages.

SUMMARY

This report analyzes the impact of the timing, spacing, and number of children on a married woman's wage growth over her life cycle. The deterrent effect of child rearing on the extent of women's labor market work has been well documented. To date, however, there have been no attempts to quantify the effect of these work reductions on subsequent growth in women's wages. It is likely that such reductions are an important explanation for the observed gap between wages of prime age men and women.

The data used for the analysis are taken from the Michigan Panel Study of Income Dynamics (IDP). This data base is unique in several respects for the purposes of this study. The ninth wave (administered in spring 1976) contains information on both prior labor market work histories and the dates of birth of all children borne by the woman. No other data base available provides this information in such detail.

The empirical approach taken in this report uses a single cross-section (1976) of the IDP data. This approach proceeds in two stages. In the first stage the amount of time the mother spends in the labor force at each point in her life cycle is related to the number and age composition of her children. The second stage consists of estimating the effect of actual accumulated labor market experience on wages over the life cycle. The two stages are linked by translating estimates of the quantity of labor market work in a given year into corresponding estimates of accumulated work experience at each point in the life cycle. This link provides the mechanism by which the impact of child rearing on wage rates is determined.

This approach permits estimates to be made of the impact of the timing, spacing, and number of births on the mother's pattern and amount of lifetime market work and the pattern of growth in her wage rate. The principal findings of this report are as follows:

The number of children a woman rears has an important deterrent effect on her accumulated work experience. A woman who has one child will work, on average, about 2.5 years less over her lifetime than a woman with no children. A second child results in an additional two years withdrawn from the labor market, and a third one additional year.

Larger birth intervals (holding constant number and timing of children) are associated with larger reductions in lifetime work effort. The magnitude of this effect appears to increase with the number of children. For a woman with three children who begins her family at age 19, we estimate that an increase in the average birth interval from two years to four years reduces lifetime labor market work by about one year. An increase in the average birth interval from four years to six years reduces lifetime labor market work by an additional one-half year.

The timing of the first birth (the age at which a woman has her first child) does not appreciably affect lifetime labor market experience. Accumulated work experience is a strong determinant of market wage rates. Evaluated at the mean experience level in the data (7 years), the rate of return to an additional year of experience was estimated to be about 4 percent.

Child rearing, through its effect on accumulated work experience, has a substantial effect on wage rates. The average wage rate over the life cycle of a woman with no children is 13 percent higher than that of a woman who has three children, spaced three years apart, beginning at age 22 (the "average" woman in our sample). The maximum difference in wage rates is 22 percent and occurs at age 32.

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1. INTRODUCTION

The demographic and economic determinants of the labor supply behavior of married women have been the subject of extensive empirical analysis.[1] One of the major outcomes of these analyses has been to quantify the deterrent effect of child rearing on the extent of women's labor force participation. The results of these studies indicate that the existence of children in the home, especially young children, is the single most important factor influencing the amount of time mothers spend at market work. There have been few attempts to relate the reduction in labor market work to women's subsequent earning capacity, despite its potential empirical importance. There can be little doubt that reduced work leads to a decline in future earning capacity. Hence, child rearing, through its effect on work experience, is potentially an explanation for at least a portion of the observed gap between wages of women and of men who have otherwise identical characteristics.

The lack of empirical analysis of this issue has stemmed primarily from a lack of data on work histories of women. With the availability of the National Longitudinal Survey of Mature Women and the 9th Year of the Michigan Panel of Income Dynamics, the absence of data is no longer a problem. The major stumbling block appears to be in devising appropriate econometric techniques for gleaning information from these data.

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[1] The number of empirical studies is far too great to mention all of them. However, the most notable in this long line of research are the classic work of Mincer (1962), which laid the foundation for almost all subsequent research on the labor supply of married women, and the extensive empirical analysis of Cain (1966), and Bowen and Finegan (1969). Later conceptual advances were made by Kosters (1966) in rigorously formulating the household model of labor supply, Gronau (1973) in formulating the labor force participation decision, and Heckman (1974) in integrating the hours of work and labor force participation decisions. Other advances have been by Heckman and Willis (1976) in estimating the degree of permanency of attachment to the labor force, and Hanoch (1976) in integrating alternative measures of labor supply.

This study is an attempt to estimate the impact of child rearing on the extent of women's labor market work and to translate this impact into its ultimate effect on her market wage rate. The study uses cross-sectional data from the 1976 survey of the Michigan Panel of Income Dynamics. This data base is chosen primarily because it contains detailed information on women's prior labor market experience. Data on both full-time and part-time years of labor force activity since age 18 are available.

Two equations characterize the basic empirical framework of this report. The first equation relates the amount of time the mother spends in the labor force at each point in her life cycle to the number and age composition of her children. This equation estimates the impact of the number, timing, and spacing of her children on the quantity of labor she supplies to the market place. The second equation estimates the relationship between her accumulated labor market experience and her market wage offer. The two equations are linked by translating estimates of the quantity of labor market work in a given year into corresponding estimates of accumulated work experience at a point in the life cycle. This link provides the mechanism by which the impact of child rearing on wage rates is determined.

In this treatment children are viewed as exogenous with respect to both the labor supply decision and the market wage offer. Although this assumption is a common one in the labor supply literature, it does run counter to the implications of work on fertility. As early as 1963, Mincer argued that a prime determinant of a woman's fertility decision is her market wage rate. As a result, fertility and labor market work are joint decisions influenced simultaneously by the woman's market wage offer. Willis' (1973) seminal work on fertility also emphasizes this point, as have many subsequent studies of fertility. To the extent that our assumption fails to hold in the data, our estimates are biased. The assumption of exogeneity, though in direct conflict with the presumption of the fertility literature, is

necessary. Without it the parameters of the labor supply and market wage equations are hopelessly underidentified.

The structure of this paper is as follows. Section 2 describes the data base used in the analyses of this report and presents a preliminary analysis. Section 3 presents estimates of the impact of the number and age of children on women's work behavior. Life-cycle labor supply profiles are derived. The profiles are translated into accumulated labor force experience measures, which, in turn, are related to various configurations of the number, timing, and spacing of births over the life cycle. Section 4 estimates the wage-labor force experience relation. Statistical issues, such as simultaneity between experience and wages and selectivity bias, are addressed. Finally, the estimates of Sections 2 and 3 are brought together to derive wage depreciation estimates resulting from time withdrawn from the labor force due to child bearing. Our conclusions are summarized in Section 5.

2. THE DATA

2.1 THE DATA BASE

The data selected for use in this study are taken from the Income Dynamics Panel (IDP). The Panel Study of Family Income Dynamics is an ongoing longitudinal survey of approximately 5000 families that is being conducted by the Survey Research Center, University of Michigan. Surveying began in 1968 and the same families have been resurveyed annually.

Several factors make this data base particularly well suited for studying the labor force behavior of women. The ninth wave (administered in the spring of 1976) included for the first time a questionnaire to be completed by wives. In addition to asking about basic demographic and recent labor force behavior, the questionnaire elicited crucial information about fertility and retrospective work history: in particular, the dates of birth of children and years of labor market experience.

The availability of children's dates of birth allows us to construct variables that measure the number, timing, and spacing of children. Several data sets that might otherwise be suitable for a study of this type are deficient with respect to information on children. The best of these, the National Longitudinal Survey of Mature Women (NLS), contains ages of all children in the home and the woman's age when her first child was born. These data are sufficient to construct the final variables used in our analysis, but the lack of information on children not living at home precluded the possibility of experimenting with more general specifications. Other large cross-sectional data files (Public Use Sample, Current Population Survey) typically contain the number of children in various age categories but not the detailed age information required.[1] The

[1]The IDP data are not perfect with respect to information on children. In particular, it is not possible to distinguish children living at home from those who are not. To a large extent the child's age does this automatically, but the possibility that some children are deceased or living with a previous spouse remains.

ninth year of the Income Dynamics Panel is unique in the amount of information provided on previous work experience of women. Three related pieces of data were collected: total years since age 18 in which the woman worked full time and part time, respectively, and of the part-time years, the usual proportion of time worked. This detailed information permits us to construct variables with which we can examine whether full-time and part-time attachment to the labor force have differential effects on subsequent wage rates.[1]

A working file was constructed from the 1976 survey (containing labor supply information referring to 1975) on white, married-spouse-present families whose wives were interviewed. These selection criteria reduced the sample from the original 5862 observations to 2394. In addition to further exclusions because of missing data for key variables, families whose wives were permanently disabled or students at the time of the survey and those whose husband or wife was self-employed were excluded. The latter restriction is due to the problems associated with separating reported income into returns to capital and returns to labor for those who are self-employed. The final sample consists of 1576 observations.

2.2 VARIABLE DEFINITIONS AND DESCRIPTIVE STATISTICS

Table 1 presents means and standard deviations of all variables used in this analysis stratified by whether or not a woman worked in 1975. Certain differences between means for workers and nonworkers stand out immediately and are noteworthy: Husband's 1975 earnings are greater for nonworkers, which is consistent with the wife's leisure being a normal good, and is also to be expected if husband's and wife's times are substitutes in household production. Working wives are somewhat more highly educated than nonworking wives, which

[1]The NLS data provide the number of years in which a woman worked 6 months or more, and therefore does not permit testing for differential effects of full-time and part-time work experience, and moreover, if the extent of work effort within each year does matter, then an unbiased estimate of this effect cannot be obtained.

Table 1

MEANS AND STANDARD DEVIATIONS BY LABOR FORCE STATUS IN 1975

(Standard deviations in parentheses)

Variable	Wife Worked in 1975	Wife Did Not Work in 1975	Total Sample
Husband's earnings in 1975[a]	11,640 (6682)	12,792 (7984)	12,404 (7184)
Age of wife[a]	33.64 (11.35)	36.61 (12.53)	34.68 (11.86)
Years of schooling completed[a]	12.65 (2.25)	11.54 (2.22)	12.26 (2.30)
Years of full-time experience[a]	7.20 (7.69)	4.59 (5.95)	6.28 (7.24)
Years of part-time experience	1.85 (3.51)	.872 (2.12)	1.51 (3.14)
Extent of part-time experience	.858 (1.79)	.423 (1.20)	.705 (1.62)
Average hourly 1975 wage	3.83 (2.28)	--	--
Hours worked in 1975	1323.8 (676.8)	0.0 (0.0)	859.27 (834.5)
Residence in South	.320 (.466)	.315 (.464)	.318 (.466)
Residence in city of 100,000 or greater population	.529 (.499)	.530 (.499)	.529 (.499)
Number of children ever born[a]	1.84 (1.74)	2.48 (1.78)	2.07 (1.78)
No children[a]	.279 (.449	.089 (.285)	.212 (.409)
Child born in 1975[a]	.081 (.273)	.125 (.330)	.0964 (.295)
Age of youngest child[b]	10.17 (8.31)	9.61 (9.41)	9.94 (8.77)
Age of wife at birth of first child[b]	21.83 (4.13)	21.92 (4.14)	21.87 (4.13)
Age of second youngest child[c]	14.90 (8.40)	14.64 (9.73)	14.79 (8.98)
Number of observations	1023	553	1576

[a]Means of workers and nonworkers are significantly different at the .05 level.

[b]Conditional on having one or more children (sample consists of 738 workers and 504 nonworkers).

[c]Conditional on having two or more children (sample consists of 532 workers and 388 nonworkers).

is not necessarily implied by theory but is consistent with previous findings (Bowen and Finegan (1969), Mincer (1962)).[1] Working wives also have fewer children and are three times more likely to have no children. Women who did not work in 1975 are half as likely to have had a child born during 1975. All of these findings are consistent with children having a deterrent effect on labor force participation and bode well for the analysis to follow.

At this point, some mention should be made of an ambiguity that arises in the handling of the children variables. A child born in 1975 can be expected to affect labor force behavior in that year differently depending on when in the year it is born. Although the month of birth is available in the data, we found that despite attempting several alternative specifications, a simple dummy variable indicating the birth of a child in 1975 produces the best results in the hours of work estimates reported in Section 3.[2]

The differences between the two groups of women in the proportion giving birth in 1975, as well as the finding that the mean ages of the youngest and second youngest children are lower for nonworkers (although not significantly so), are all the more noteworthy because the mean age of the workers is about 3 years less than that of the nonworkers.

Not surprisingly, women who worked in 1975 had considerably more previous labor market experience than those who did not work. The definitions of the experience variables are as follows: full-time years of experience is the number of years since age 18 that were worked full time. Part-time years of experience is the number of years since age 18 in which the woman worked but worked less than full time.

---

[1] Since schooling presumably increases both market and nonmarket productivity, its effect on participation is ambiguous.

[2] Separate dummy variables were used to represent month of birth, but revealed no clean-cut pattern. It is possible that a more exhaustive search would have turned up a better functional form, but such an effort was deemed too costly.

Extent of part-time experience is part-time years of experience times the proportion of time usually worked in those years. Total years of experience is the sum of full-time years and extent of part-time experience and is our richest measure of amount of previous labor force experience.[1] Workers in 1975 had previously accumulated 61 percent more total experience than had nonworkers. Expressed somewhat differently, workers had accumulated 89 percent more total experience than nonworkers when experience is expressed as a proportion of the time since age 18.

Table 2 relates number of children and age of the youngest child to degree of labor market participation in 1975. Several patterns of relationships can be discerned. There is a generally rising level of participation with respect to age of youngest child and a negative relationship to number of children, suggesting the presence of separately estimable effects of numbers and ages of children. Also worthy of note is the relatively high degree of participation by women with young children.

Table 3 presents full-time experience and extent of part-time experience (expressed as a proportion of time since age 18) by number of children. Full-time years of experience is seen to decline with the number of children. Breaking accumulated experience down by the age of the youngest child and by mean birth intervals did not reveal any systematic relationship, suggesting that 1975 work effort may be a more suitable variable with which to examine spacing effects.

Table 4 presents statistics on average hourly earnings among 1975 workers by total years of experience. It is interesting to note how quickly wage rates level off as experience increases. Rather than interpret this as a life-cycle pattern, one must bear in mind the likelihood that there are other factors that affect wages that are correlated with experience. For example, women with less experience (on average, younger cohorts) may be more highly educated, and the education they do have, being more recently attained, may be more relevant to current market productivity than that of older cohorts.

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[1] All experience measures have been adjusted to remove 1975 work experience. Hence they refer to experience accumulated prior to 1975.

Table 2

PROPORTION OF TIME SPENT AT WORK BY NUMBER OF CHILDREN AND AGE OF YOUNGEST CHILD[a]

(Cell sizes in parentheses)

Age of Youngest Child	Number of Children[b] 1	2	3+	All Women
Child born in 1975	.29 (70)	.27 (56)	.18 (26)	.26 (152)
1-2 years	.25 (49)	.24 (43)	.27 (24)	.25 (116)
2-3 years	.34 (36)	.12 (28)	.40 (21)	.28 (85)
3-4 years	.54 (25)	.41 (26)	.20 (22)	.39 (73)
> 4 years	.63 (142)	.49 (235)	.45 (439)	.49 (816)
All ages	.46 (322)	.40 (388)	.42 (532)	.42 (1242)

[a]The dependent variable is hours worked in 1975 divided by 1750 hours (a conventional definition of a full-time year).

[b]For women with no children the proportion of time spent at work in 1975 is .74 (334 observations).

Table 3

ACCUMULATED LABOR FORCE EXPERIENCE BY NUMBER OF CHILDREN[a]

Variable	Number of Children 0	1	2	3	4+
Full-time years	.46	.48	.39	.30	.25
Extent of part-time years	.07	.04	.05	.04	.04
Cell size	334	322	388	249	283

[a]The dependent variables, full-time years of experience and extent of part-time experience (as defined previously), are expressed as proportions of the total time since age 18.

Table 4

MEAN AVERAGE HOURLY EARNINGS BY TOTAL YEARS OF EXPERIENCE

Item	Years of Experience 0	1-3	4-6	7-9	10-12	13-18	19-24	>24
Mean	2.18	3.09	3.90	4.72	4.37	4.19	4.30	4.33
Standard Deviation	1.36	1.82	2.20	2.54	2.68	2.27	2.16	2.37
Number	52	272	246	157	80	98	67	51

3. CHILD BEARING AND ACCUMULATED WORK EXPERIENCE

3.1 INTRODUCTION

The approach to estimating the impact of childbearing on the mother's accumulated work experience adopted in this report consists of two stages. The first stage estimates the expected amount of labor market work at each point in the woman's life cycle. The measure of the amount of labor market work in a given year is annual hours of work. The primary focus of the analysis of this stage is on the impact of the number and age composition of children. The second stage translates these estimates into estimates of accumulated work experience to obtain estimates of the effect of the number, timing, and spacing of children on the wife's lifetime labor supply.

This study uses cross-sectional data to estimate the relationship between childbearing and accumulated work experience. Estimates of life-cycle behavioral relationships from cross-section data require the strong assumption that the underlying behavioral parameters across women of different age cohorts are identical. Herein this amounts to assuming that labor supply and wage equation parameters across women of different ages are the same. There is some evidence that cohort differences in labor supply and wage parameters are important and, hence, our assumption is not valid. Age-specific labor force participation rates among women were considerably higher in 1970 than in 1960, especially among young women. If the rising age-specific participation rates reflect cohort differences, our estimated age-related parameters will measure a combination of life-cycle and cohort effects. However, this evidence is not conclusive. Rising age-specific female labor force participation rates may result from rising demands for female labor, rather than from changes in the underlying labor supply parameters. Hence, it is not clear whether our assumption of no cohort differences is inappropriate. But, it is important to note our assumption and to remark that life-cycle and cohort effects cannot be distinguished in cross-sectional data.

Panel data on women in our sample are available, but the budget of this study did not permit exploiting them.

3.2 THE BASIC STATISTICAL FRAMEWORK

The first step in our analysis is to estimate the relationship between the number of hours a woman can be expected to work in any given year of her life cycle as a function of the number and age of her children, and various other demographic and economic characteristics that are presumed to affect her labor supply decision. The conceptual model that underlies our estimation procedure draws from earlier research in labor supply, most notably that of Gronau (1973b) and Heckman (1974).

In this model, the woman's decision to participate in the labor force and her choice of hours of work are made simultaneously. Both of these decisions are cast in terms of two variables: a market wage offer (W) and a reservation wage (W_R). The reservation wage is the lowest wage at which the woman would be willing to enter the labor force. Thus,

$$(3.1) \qquad h = \max \{0, g(W-W_R)\}$$

where $g(\cdot)$ is greater than zero and, assuming the labor supply function is upward rising, $g'(\cdot) > 0$.

The hours of work equation indicates that if the market wage offer exceeds the reservation wage, the woman participates in the labor force and her hours of work are positive. If the reservation wage exceeds the market wage, she does not participate in the labor force and her hours of work are zero.

The woman's demographic and economic characteristics as well as the number and age composition of her children enter the analysis through their influence on the market wage offer she faces and/or her reservation wage. By substituting variables representing these characteristics and the composition of her family, the hours of work equation may be expressed as a function of these variables. To

illustrate, assume that the market wage and reservation wage equations are linear functions of their respective determinants. Denoting the determinants of the market wage by the vector X_1 and the determinants of the reservation wage by X_2, we may express these wages as

$$(3.2) \qquad W = B_1'X_1 + \varepsilon_1$$

$$(3.3) \qquad W_R = B_2'X_2 + \varepsilon_2$$

where ε_1 and ε_2 measure the effect of unobservable variables on the market and reservation wages, respectively. B_1 and B_2 are unknown vectors of parameters that measure the influence of X_1 and X_2, respectively.

Assuming further that $g(\cdot)$ in equation 3.1 is linear and substituting equations 3.2 and 3.3 into the hours of work equation, we may express the latter equation as

$$(3.4) \qquad h = \max \{0, \gamma[B_1'X_1 - B_2'X_2 + \varepsilon_1 - \varepsilon_2]\}$$

Equation 3.4 is termed a reduced form hours of work equation. Once the distribution of the disturbances is specified, the reduced form parameters, $\pi_1 = \gamma B_1'$ and $\pi_2 = -\gamma B'_2$, can be estimated with maximum likelihood.

In this study we have used Tobit analysis (Tobin, 1958) to estimate the reduced from parameters. The Tobit estimation procedure assumes that the disturbance terms, ε_1 and ε_2, are normally distributed with mean zero, constant variance, and constance covariance. In the Tobit formulation the expected number of hours a woman will work in a given year is related to her characteristics by the equation

$$(3.5) \qquad E(h|X_1,X_2) \equiv P\{h > 0\}\, E(h|X_1,X_2,h > 0) =$$

$$P\{h > 0\}\, I + \sigma_u f\left(\frac{I}{\sigma_u}\right)$$

where

$$I = \pi_1 X_1 + \pi_2 X_2$$

$$\sigma_u = (\sigma^2_{\varepsilon_1} + \sigma^2_{\varepsilon_2} - 2\sigma_{\varepsilon_1\varepsilon_2})^{1/2}$$

The term $f(\frac{I}{\sigma_u})$ is the normal density function evaluated at I/σ_u. The term $P\{h > 0\}$ is the probability that the woman works, which is given by the equation[1]

$$P\{h > 0\} = \int_{-\frac{I}{\sigma_u}}^{\infty} \frac{1}{\sqrt{2\pi}} e^{-1/2t^2} dt$$

3.3 SPECIFICATION OF THE HOURS OF WORK EQUATION

In the literature on married women's labor supply behavior, there is fairly wide agreement on the set of variables that affect that woman's hours of work and the probability that she works. There is, however, less agreement on whether these variables affect the woman's labor supply through their effect on market wages, reservation wages, or both. For example, studies by Schultz (1975), Heckman (1974), and Cogan (1975) all use education as an explanatory variable in reduced form hours of work equations. Schultz argues that the education effect operates only through its impact on market wages. Heckman and Cogan, on the other hand, following the work of Michael (1973), argue that education affects labor supply through its effect on both market and reservation wages.

Furthermore, even though there may be a consensus on whether a particular variable influences the market or reservation wage, there is virtually no agreement on the functional form of the variable.

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[1] A more complete discussion of the Tobit model is provided in Tobin (1958) and Amemiya (1973).

Children in the home are a prime example. There is virtual unanimity in the belief that children influence the woman's reservation wage but have no effect on the market wage offers, except to the extent that children may serve as a proxy for prior labor force experience. However, some studies (Heckman, 1974) assume that only preschool children affect the wife's reservation wage. Others, such as Cain (1966), Bowen and Finegan (1969), Gronau (1973a), and Hanoch (forthcoming), assume that the total number of children as well as their age composition affects the wife's reservation wage. Moreover, in each of the latter four studies a different set of variables is used to capture the effect of the number and age composition of children.

In specifying the reduced form hours of work equation, we dichotomize the right-hand-side variables into two sets: variables measuring the number and age composition of the woman's children and variables measuring her other demographic and economic characteristics.

The second set includes a rather conventional group of explanatory variables. They are: the level of the husband's earnings, the wife's age, a dummy variable indicating residence in the South, a dummy variable indicating residence in an urban area of population over 100,000, and the wife's level of education. The theoretical and/or empirical importance of each of these variables in affecting married women's labor supply is well known. Hence, they will only be briefly discussed.

The husband's earnings affect the wife's labor supply through its effect on her reservation wage. If the wife's nonmarket time is a normal good, a rise in the level of her husband's earnings will increase the quantity of nonmarket time demanded and, hence, raise the minimum wage rate necessary to just induce her to enter the labor market. The wife's age affects her labor supply through its effect on both her market wage offer and her reservation wage. Age-market wage profiles of married women tend to rise at young ages, and then decline. Therefore, the partial effect of age which operates through the market wage should exhibit the familiar inverted U-shaped pattern.[1]

[1] A more detailed discussion of the effect of age on the wage offer is deferred until Section 4.

When the wife's labor supply decision is viewed in a life-cycle context, her reservation wage would be expected to vary with her age, though its expected pattern is not clear.[1] The southern residence dummy variable is included to capture the fact that wages of women in the southern states of the U.S. are lower than those in other regions of the country (Smith, 1978). The urban residence dummy is included to capture the fact that wages paid to women in urban areas tend to be higher, ceteris paribus, than wages paid to women in rural areas (Fuchs, 1971; Cogan, 1977). The wife's level of education is presumed to operate both through her market wage and her reservation wage. Education is viewed as an investment in human capital, raising both market and nonmarket productivity (Michael, 1973). Higher market productivity is reflected in higher wage rates, which tend to increase the wife's hours of work. Higher nonmarket productivity alters the wife's reservation wage. Whether it raises or lowers her reservation wage is, however, theoretically ambiguous.[2]

In specifying the functional form of the hours of work equation, each of the variables in this set was entered linearly. The focus of our initial efforts at arriving at a suitable functional form was on various ways of measuring the age composition and number of children. The many specifications estimated differed only in the variables used to characterize the spacing, timing, and number of children. Initially, we included separate dummy variables for the ages of the youngest, second youngest, and third youngest child; these interact, of course, with dummy variables indicating that the women had such a child. These dummy variables allow the effects of ages of children to affect participation probabilities in any arbitrary manner for a woman who has at most three children. We also included dummary variables for each number of children the woman could have had. This allows for the

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[1] See, for instance, Heckman (forthcoming).

[2] Michael (1973) rigorously demonstrates this point.

possibility that it is the presence of a child, rather than its age, that affects labor supply. Finally, we included dummy variables representing the age of the wife when she had her first child. The four sets of age-related dummy variables were used to trace out a time path of coefficients. A functional form corresponding to each time path was chosen and used to obtain the final estimating equation.

The coefficients on the dummy variables indicating the presence of a youngest child of a given age showed, as expected, that the younger the child, the greater its deterrent effect on the mother's hours of work. The effect of the presence of a child one year old reduced the mother's labor supply by approximately 1050 hours per year. The deterrent effect of the youngest child was found to decline at a decreasing rate as the child ages. When the child reaches 8 years of age its negative impact on the mother's labor supply was estimated to be about 160 hours per year. At age 10 the deterrent effect of the youngest child fell to about 90 hours per year and was not statistically significant. The functional form of the age of the youngest child chosen to correspond to this time path is the inverse of the age of the youngest child. This form of the variable permits the youngest child to have its largest negative impact at age 1 and declines at a decreasing rate as the child ages.[1]

The coefficients on the dummy variables indicating the presence of a second youngest child of a given age showed that the second youngest child had a significant deterrent effect on the mother's labor supply up to the point where the child reaches age 13. Thereafter, its effect was not statistically significant. The coefficients displayed no discernible pattern from age 1 to age 5, but systematically declined in absolute value thereafter. We chose to constrain the effect of the age of the second youngest child to be constant over the age range 0 to 6 and to decline thereafter. The variable included to capture this decline was the inverse of the age of the child.

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[1]The natural logarithm of the age of the youngest child would also trace out the same path. It was used but did not fit the data as well as the inverse of the age of the youngest child.

The coefficients on the dummy variables indicating the presence of a third youngest child of a given age were never statistically significant and showed no consistent pattern. It appears that once the ages of the two youngest childrn are controlled for, the age of the third youngest child does not matter and was therefore not included in the final specification.

The coefficients on the dummy variables of the last set measure the age of the mother when she gave birth to her first child. They were never statistically significant and showed no pattern. Hence, any measure of the age of the mother at birth of her first child were omitted from the set of regressors used in the final specification.

3.4 HOURS OF WORK ESTIMATES

Table 5 gives parameter estimates of the hours of work equation. Since these estimates were obtained after considerable experimentation with the children variables, the reported standard errors should be regarded with caution. Column one gives the Tobit index coefficients, which are estimates of the B's in equation 3.4. Column three reports changes in the expected value of work for unit changes in the value of the right-hand-side variable. These estimates are computed by differentiating the expected hours of work function (equation 3.5) and evaluating these derivatives at the sample means.

The coefficients of the demographic and economic variables other than those measuring the number and age of children all have their expected signs and are quantitatively similar to estimates reported in other studies. The husband's earnings exert a negative effect on the wife's labor supply. This estimate conforms to that obtained by Heckman (forthcoming) of -16.3 hours per year using the 1967 National Longitudinal Survey of Married Women. Cogan (1977), using the same data base as Heckman, estimated a virtually identical effect to that obtained in this study: -22.4 hours per year. Hanoch (forthcoming) estimates a much larger effect of -92.4 hours per year using the 1967 Survey of Economic Opportunity data. The age coefficient is negative, consistent with the results of Greenberg and Hosek (1976)

Table 5

HOURS OF WORK PARAMETER ESTIMATES

Variable	Tobit Index Coefficient	Standard Error	Expected Value Coefficient
Husband's earnings ($1000)	-22.6	4.53	-16.0
Wife's age	-17.55	3.62	-12.43
Residence in the South	134.40	64.61	95.16
Residence in urban area	-23.41	60.49	-16.57
Wife's education	99.6	14.32	70.53
(Age of youngest child)$^{-1}$	-1170.71	253.56	-828.98
(Age of second youngest child)$^{-1}$	-1587.63	780.93	-1124.20
Number of children aged 1 to 6	-228.24	73.13	-161.62
Child born in 1975 dummy	-942.43	115.96	-667.33
Number of chldren aged 6 and over	9.16	24.36	6.49
Intercept	577.24	227.53	

and Schultz (1976). The residence in the South dummy variable is positive, indicating that southern women work on average three more weeks per year than women residing in the rest of the country. This is somewhat curious in that real wages of women tend to be lower in the South than in other regions of the country. The urban residence dummy is small and not significant, suggesting no urban-rural hours of work differential despite the tendency for urban wages to be higher than rural wages. The wife's education is positive, consistent with virtually all empirical analyses of the labor supply behavior of married women.

The effects of the children variables are large and statistically important. A good idea of the magnitude of the effects can be obtained from Table 6, which computes the effect of the youngest and second youngest children at various ages. The calculations of the effect of the youngest child are made assuming that the wife has no other children. As column 1 indicates, the reduction in work declines rather rapidly the older the youngest child. The reduction in work for a working woman if the youngest child is one year old is about 35 weeks per year, whereas the reduction associated with a youngest child aged 7 is only 4 weeks per year. The calculations of the effect of the age of the second youngest child are made assuming that the woman has only two children and that the age of the youngest child is one year. The reduction in hours of work associated with variations in the age of the second youngest child, given the youngest child is one year old, is somewhat smaller than that associated with variations in the age of the youngest child.

A more complete picture of the effect of children on the labor supply behavior of women can be obtained from Figure 1. Figure 1 depicts the expected work reductions, expressed as a fraction of a full-time year of work foregone, at each point in the life cycle associated with three family sizes. This measure of work reduction is the difference between what she would have worked if she had no

Table 6

PARTIAL EFFECTS OF AGE OF YOUNGEST AND SECOND YOUNGEST CHILD

Youngest Child		Second Youngest Child	
Age of Child	Reduction in Work Hours per Year	Age of Child	Reduction in Work Hours per Year
1	1398	6	1654
3	618	7	1616
5	462	9	1565
7	158	11	1533
9	121	13	1511
11	97	15	1495
13	81	17	1482

children and what she works given her children. These profiles were computed from the estimated expected hours of work function (equation 3.5).[1]

The solid line in the figure depicts the life-cycle path of work reduction for a woman who has the average characteristics in the data except that she has a child at age 19. The year in which she bears her child she reduces her market work effort by 40 percent of a year. In her first full year with her child, age 20, she reduces her market work by 53 percent. As her child grows to age 5 she returns to market work at a rapid, but decreasing rate. As the woman reaches 24 years of age, her child is 5 years old and the reduction in her labor supply is 23 percent of a year. The next year, as her child enters elementary school, she further increases her labor supply. Thereafter, from age 25 to age 37, the quantity of labor she supplies to the market increases

[1]Full-time work is assumed to be 1750 hours per year. This is 35 hours per week times 50 weeks per year.

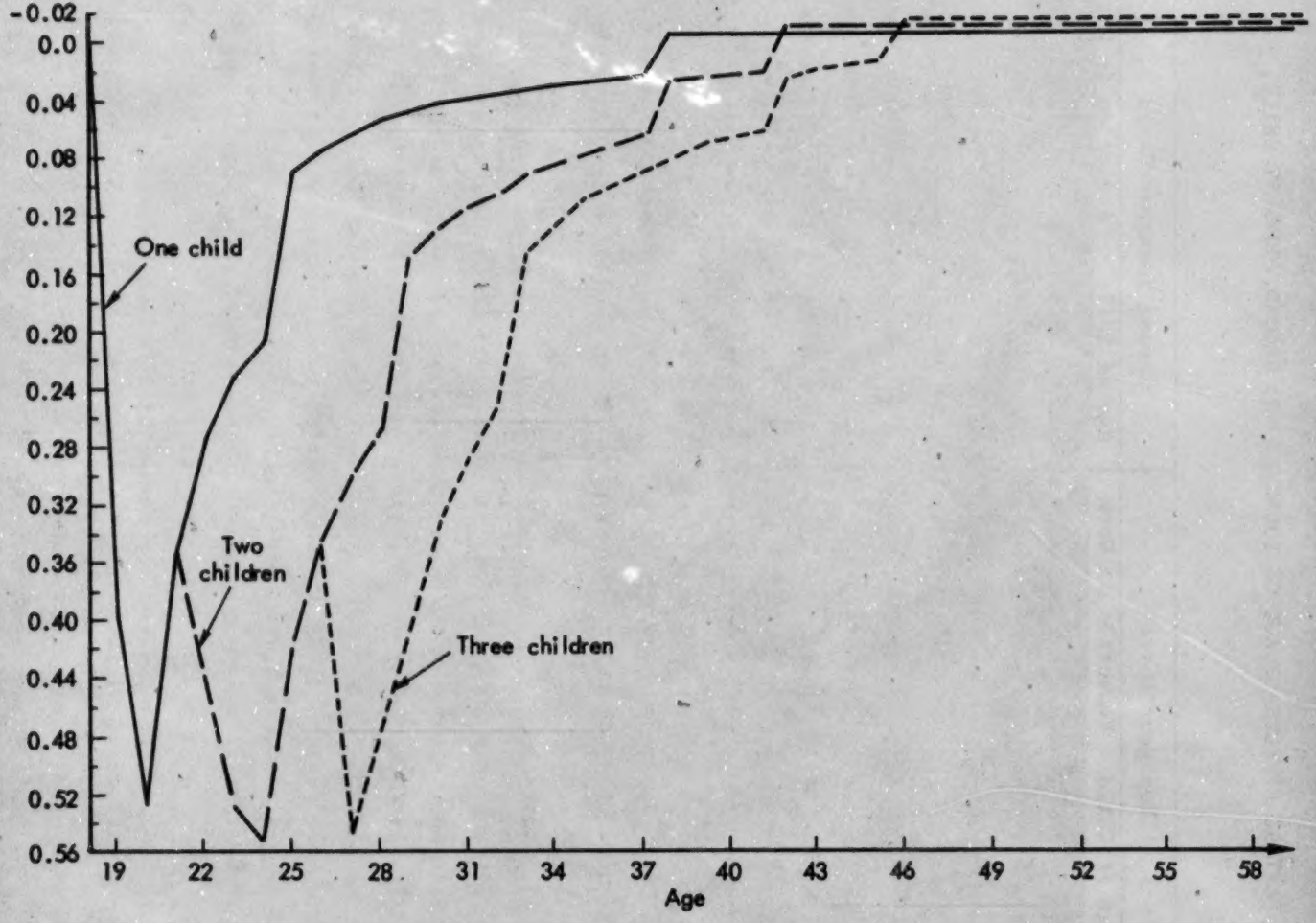

Fig. 1—Expected reductions in market work

gradually and is only slightly below what she would have worked in the absence of having the child. At age 38, as her child presumably leaves home, she returns to what she would have worked if she had no children.[1]

The profile described by the dashed line represents the expected life-cycle reduction in market work for a woman who bears two children; one when she is age 19, the other when she is age 23. Her life-cycle profile is identical to that of a woman who bore only one child up to the point where she bears her second child.[2] The year she bears her second child the amount of her market work declines by 53 percent of full-time work, and then declines by an additional 3 percent the first year she bears the child. Thereafter, her profile rises much in the same manner as a woman who bore only one child--rapidly as the oldest child enters elementary school and then leveling off after the second child reaches school age. The woman with two children returns to work as much as she would have in the absence of any children about four years later than the woman with only one child.

The profile described by the third line represents a woman who has three children--the first at age 19, the second at age 23, and the third at age 27. The reduction in labor market work during her child bearing years is sizeable. Her re-entry rate is quite rapid after she completes her fertility. By age 35 she spends almost as much time in the labor force as a woman with only two children. This presumably reflects the availability of the oldest child as a baby-sitter or as a helper around the house. The oldest child is 16 when the mother reaches age 35. After age 35 her re-entry rate is slower than that of a woman with two children. This probably reflects the deterrent effect of the youngest child and the loss of the services

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[1]This result may seem surprising since we would expect her market wage offer to be lower because her accumulated experience is lower. However, we found no evidence that married women who have children over 18 work less than those who never had a child.

[2]We included an interaction term between the age of the first child and the existence of more than one child to capture the possibility that re-entry rates into the labor force may depend upon whether or not the family has been completed. The interaction term was not significant.

of oldest child, as this latter child leaves home. The woman with three children does not return to what she would have worked in the absence of any children until age 46.

3.5 EFFECTS OF THE NUMBER, TIMING, AND SPACING OF CHILDREN ON ACCUMULATED WORK EXPERIENCE

Cross-section regression estimates of the effect of the number and age composition of children on labor supply variables alone cannot possibly provide information on the effects of the timing and spacing of births on the mother's labor supply behavior. This is because variations in the timing and spacing of births necessarily result in variations in the ages of the children. The lack of any independent variation among these variables precludes identification of the independent effects of each. For example, suppose we observe two women 30 years of age who each have one child, have completed their families, and have otherwise identical characteristics. Imagine that one woman bore her child when she was 20 years old and is currently working 2000 hours. The other woman, let us suppose, bore her child when she was 25 years old and is currently working 1000 hours. Is the difference of 1000 hours attributable to the fact that one woman began her family earlier than the other, or is the difference due to the fact that one woman has a child 10 years old in the home while the other has a child 5 years old in the home? We cannot say, without arguing that one of the two effects is zero.

The issue of the relation between the timing and spacing of births and the mother's labor supply only makes sense when the wife's labor supply is viewed within the context of her lifetime labor supply. A summary measure of her lifetime labor supply is the amount of accumulated work experience. In this subsection we translate the cross-section hours of work estimates into lifetime accumulated labor market experience.

One rather obvious measure of the amount of accumulated work experience is the sum of the proportion of each year spent at market work from age at completion of schooling to age at retirement. Unfortunately, we do not have data on either of these ages. Without

these ages we have chosen to measure accumulated work experience as the proportion of years spent at market work from age 19 (for a woman with 12 years of schooling) to age 60.

This measure is not without its pitfalls. The age at which a woman chooses to retire may be influenced by the number, timing, and spacing of her children. If there is a relationship and some women retire after age 60, our measure is biased.

Our measure of accumulated work experience is computed by simply summing the expected proportion of each year spent at market work. The expected proportion of time worked is calculated as the predicted hours of work from the regression estimates divided by a measure of full-time work, 1750 hours. Computing accumulated experience in this manner requires the strong assumption that intertemporal variations in hours of work are not state dependent, i.e., the amount of time a woman spends at market work in a given year is not a determinant of the amount of time she works in a subsequent year. Recent work by Heckman (forthcoming) has indicated that a related dimension of labor supply, labor force participation rates of women, is determined by a state dependent process. To the extent that there is state dependence, our estimates of the effect of children on accumulated work experience are biased toward zero. However, we must live with this possible bias. Incorporating state dependence into the analysis requires the use of maximum likelihood estimators, the cost of which exceeds the budget of this study.

To determine whether the timing of the birth of the first child has any impact of the mother's accumulated work experience, we considered three possible ages of the mother at first birth and three alternative total numbers of children. The three ages at first birth are ages 19, 23, and 27. The total number of children are one, two, and three. The spacing between the children in the latter two cases is two years. The results indicate that the timing of the first birth does not appreciably affect the extent of the mother's market work over her life cycle. The largest difference is between two women

who have three children, one of whom starts her family at age 19, the other at age 27. The eight year differential in timing between these two women results in only a 0.3 year differential in labor market work.

There is some evidence that the spacing between births affects the amount of lifetime labor market activity. Table 7 reports spacing results for women with two and three children.

Table 7

BIRTH SPACING AND LIFETIME LABOR MARKET WORK

Two Child Family		Three Child Family	
Age of Wife at Birth of Children	Lifetime Work (yr)	Age of Wife at Birth of Children	Lifetime Work (yr)
19-21	19.03	19-21-23	17.9
19-23	18.70	19-23-27	17.0
19-25	18.6	19-25-31	16.5

The table indicates that for a three-child family an increase in the average birth interval between children from two to four years reduces lifetime labor market work by one year. A further increase in the average interval from four years to six years reduces lifetime labor market work by an additional one-half year. For a two-child family the pattern is the same, but the magnitudes are much smaller.

The most important effect of child rearing on lifetime labor market work operates through the number of children the wife bears. Table 8 presents (by selected spacing patterns) the marginal reduction in lifetime work effort associated with increasing the number of children (for women who had their first child at age 19).

Table 8

MARGINAL REDUCTION IN LIFETIME WORK

(Years)

Number of Children	2 Year Birth Interval	4 Year Birth Interval	6 Year Birth Interval
2	2.0	2.3	2.4
3	1.1	1.7	2.1

Women who bear no children work, on average, 23.5 of the 42 years from age 19 to age 61. A woman with one child works 2.5 years less over her lifetime. A woman with two children works an additional 2.0 to 2.4 fewer years depending on her spacing pattern, and a woman with three children works an additional 1.1 to 2.1 fewer years over her lifetime. The life-cycle expected work reduction profiles in Figure 1 provide a good picture of how reduced work is distributed over the life cycle. Almost all the reduction in market work for a woman with one child occurs within the first six years of the birth of the first child; for a woman with two children spaced two years apart, it occurs within the first eight years after the birth of the first child; and for a woman with three children spaced two years apart, it occurs within the first sixteen years after the birth of the first child.

Note in Table 8 that the marginal reduction in lifetime labor supply declines with the number of children. Several explanations are consistent with this pattern. One is simply learning behavior--that is, as a woman has additional children she learns to more

efficiently allocate her time and other household inputs, thereby possibly reducing the labor market time withdrawn.[1]

Another explanation is that there exist economies of scale in the production of children. It seems intuitively reasonable that increasing the number of young children at home from one to two or from two to three will increase the various demands made on the mother's time less than proportionately. Applying standard production theory in the present context, increasing returns to scale can be due to a number of factors including "excess capacity" or indivisibilities in the woman's nonmarket time allocations, the potential for an increased degree of specialization (especially between husband and wife) in child rearing methods, fixed costs in child-care-related activities such as meal preparation and transportation to school, and quantity discounts in baby sitting and other day care services. By spacing children close enough together, a woman can take advantage of some or all of these sources of increasing returns to scale to minimize the cost of child rearing.

To the extent that any of these phenomena are manifested in the data, we would expect that women who space their children further apart would exhibit marginal lifetime work reduction profiles that decline less rapidly, either because information learned while rearing earlier children depreciates as the interval between births increases, or because scale economies become less important as birth intervals increase (for example, children of similar ages are more likely to go to the same school, eat the same food, have the same books read to them, etc.). In other words, the marginal cost of providing certain services for a child will be lower if the same or similar services are already being provided for another child. A glance at Table 8 reveals that, indeed, marginal lifetime work reduction declines less rapidly as birth intervals increase.

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[1]Again, see Michael (1973) on the relation of nonmarket productivity and reservation wages.

4. EXPERIENCE AND WAGES

4.1 INTRODUCTION

In this section we discuss estimation of the effect of experience on the wage offer. There are two potential statistical problems in estimating wage offer equations for women. The first is that prior accumulated labor market experience is likely to be endogenous with respect to current wages. If there are permanent differences in wage offers due to unobservable factors, such as market-related ability, then we would expect to observe a positive correlation between wages and labor market experience independent of any causal effect of experience on wages. Estimation of the wage equation ignoring this factor would result in a biased estimate of the true causal effect of experience on wages. The estimated effect of experience would measure movements across wage profiles of women with differential abilities, as well as its true causal effect on the wage of a given woman. The second problem is that wages are not observed for non-working women. This gives rise to the possibility of a "selectivity bias" if ordinary least squares is used on a sample of working women (Gronau, 1974). Before turning to the specification of the wage equation, it will be useful to consider each of these problems and the methods we have adopted for dealing with them.

4.2 THE ENDOGENEITY OF ACCUMULATED LABOR MARKET EXPERIENCE

Accumulated labor market experience is likely to be endogenous with respect to current wages for two reasons. One reason is reverse causality. If women anticipate the growth in wages that results from the amount of work, they may be induced to enter the labor force. Women with high growth rates are more likely to work. Since current wage rates, in part, reflect growth rates, causality would run backward in time, from current wages to prior labor force experience.

A second reason stems from the fact that levels of wage profiles are likely to differ across women. Women with high wage profiles are more likely to work at all points in the life cycle than women with

low wage profiles, assuming, of course, an upward rising lifetime labor supply function. Thus, women with high wage rates at a given point in time will have large amounts of accumulated work experience. Ordinary least squares will estimate a positive "effect" of experience on wages regardless of whether, in fact, any true causal effect exists.

The problem of endogenous labor market experience can be handled with the use of an instrumental variable. Labor market experience is a limited variable. It cannot fall below zero, nor can it exceed the number of years the woman could have worked full time. In earlier work, Cogan (forthcoming) shows that an appropriate instrumental variable is the expected value of experience. He further demonstrates that an estimate of the expected value that converges in probability to the true expected value is also an appropriate instrumental variable. Such an estimate can be obtained by the use of Two-Limit-Tobit (Rosett and Nelson, 1975) applied to a reduced form experience equation. An alternative and simpler approach, following the work of Amemiya (1973), is to approximate the expected value with ordinary least squares. This approach accounts for the inherent nonlinearity in the expected value locus by including quadratic terms for each of the right-hand-side variables and interaction terms between the right-hand-side variables.

The latter approach to obtaining an instrumental variable is adopted in this study. The instrumental variable estimates, reported in Appendix A, did not reveal any evidence of endogeneity of labor market experience. The point estimate of the effect of experience on wages was larger in the instrumental variable equation than in the ordinary least squares equation. This result is counter to what we expected. The conclusion we draw is that if there is simultaneity between current wages and accumulated work experience the instrumental variable estimates do not reveal it.[1]

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[1]Interestingly enough, Mincer and Polachek (1974) also find that the instrumental variable estimate exceeds the ordinary least squares estimate.

4.3 SAMPLE SELECTION BIAS

The fact that we do not observe wages for women who do not work creates what has become known as a "sample selection bias" problem (Gronau, 1974). An effect of the problem is that we do not observe the full wage offer distribution facing married women, only a portion of it. Ordinary least squares applied to the subsample of working women will reflect variations in means of wage distributions conditional upon work rather than *means* of the complete wage offer distribution.

To illustrate the problem of selectivity bias and to describe the method we have adopted to eliminate this bias, consider the following formulation of the process which generates wage and labor force participation.[1] The wage offer equation is given by

$$W = \beta_1' X_1 + \varepsilon_1 \tag{4.1}$$

Using our earlier notation, a woman will be observed to be working if

$$W > W_R \equiv \beta_2' X_2 + \varepsilon_2 \tag{4.2}$$

or, equivalently, if

$$u = \frac{\varepsilon_2 - \varepsilon_1}{\sigma_u} > \frac{\beta_1' X_1 - \beta_2' X_2}{\sigma_u} = \frac{I}{\sigma_u} \tag{4.3}$$

Among the subsample of working women, the expected value of wages conditional upon the values of its exogenous determinants is

$$E(W|X_1, u > \frac{-I}{\sigma_u}) = \beta_1' X_1 + E(\varepsilon_1 | u > \frac{-I}{\sigma_u}) \tag{4.5}$$

Equation 4.5 indicates that if the conditional expectation of the disturbance term (ε_1) is nonzero, a "selectivity bias problem" arises. The expected value of the term ε_1 is not zero if ε_1 and u

[1] This formulation of the problem is due to Gronau (1973b). The method of resolving it is due to Heckman (forthcoming).

are correlated; that is, if unobserved factors which affect the woman's wage offer are correlated with unobserved factors which influence the probability of her working.

Assuming ε_1 and u are normally distributed, equation 4.5 may be rewritten as

$$(4.6) \qquad E(W|X_1, u>\frac{-I}{\sigma_u}) = \beta_1'X_1 - \frac{\sigma_{\varepsilon_1 u}}{\sigma_u} \frac{f\left(\frac{I}{\sigma_u}\right)}{P\left(\frac{I}{\sigma_u}\right)}$$

where $\sigma_{\varepsilon 1u}$ is the covariance between the market wage and labor force participation equation residuals, and $f(\cdot)$ and $P(\cdot)$ are the density and distribution functions for the unit normal distribution, respectively. Using 4.6, we may rewrite the wage offer equation among the subsample of labor force participants as

$$(4.7) \qquad W = \beta_1'X_1 - \frac{\sigma_{\varepsilon_1 u}}{\sigma_u} \frac{f\left(\frac{I}{\sigma_u}\right)}{P\left(\frac{I}{\sigma_u}\right)} + v$$

where v has mean zero in the subsample. Heckman (forthcoming) has shown that if a consistent estimate of $f(\cdot)$ and $P(\cdot)$ can be obtained ordinary least squares could be applied to equation 4.7, using their ratio as a right-hand-size variable to obtain consistent estimates of β_1'. He suggests Probit analysis applied to the labor force participation decision should be used to estimate $f(\cdot)$ and $P(\cdot)$.

Heckman's approach is adopted in this study. The estimates are reported in Appendix B. The inclusion of the inverse of Mills ratio had virtually no effect on the parameter estimates. Its coefficient was statistically significant but negative, indicating the anomalous result that wages of working women are less than wages of nonworking women. There appears to be no evidence of selectivity bias in the ordinary least squares estimates. Previous work on sample selection

bias by Heckman (forthcoming) and Cogan (1977) also found little evidence of selectivity bias in women's wage equations.[1] Given our results, the issue of selectivity bias is ignored hereafter.

4.4 WAGE EQUATION PARAMETER ESTIMATES

The specification of the wage equation is

$$(4.8) \qquad L(W) = \beta_0 + \beta_1 SOUTH + \beta_2 CITY + \beta_3 EDUCATION + \beta_4 AGE + \beta_5 LEXP + \varepsilon_1$$

where L(W) is the natural logarithm of the hourly wage rate, SOUTH is a dummy variable indicating residence in the South, CITY is a dummy variable indicating residence in a city of 100,000 or more people, EDUCATION is the wife's level of education in years, AGE is her age, and LEXP is the natural logarithm of her accumulated labor market experience. Accumulated experience is defined as the number of years in which the woman worked full time plus the product of the number of years she worked part time and the percentage of time spent at work during these years.

This specification was arrived at after considerable experimentation. The focus of our experimentation was on the way in which prior experience affects wages. The results of some of this experimentation are of interest and deserve mention.

We first entered two measures of experience: the number of years of full-time work and the number of years of part-time work, as separate right-hand-side variables.[2] This was done to determine whether there is a differential return to full-time versus part-time work. As one would suspect, the estimate of the effect of full-time

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[1] In Cogan's (1977) equation, the coefficient on the inverse of Mills ratio was significant and positive, but its inclusion did not alter the estimated parameters in an appreciable manner.

[2] In each of the different specifications, the experience variable was entered both linearly and logarithmically.

work exceeded that of part-time work. The two estimates were significantly different at the 5 percent level, indicating that the amount of time spent at market work during a working year matters in determining wage growth.

We then constructed a variable measuring the extent of part-time work--the product of the number of years of part-time work and the proportion of time spent at work during these years. The wage equation was re-estimated, replacing the number of years of part-time work with this variable. In this specification, the coefficients of full-time work and the extent of part-time work were both significantly different from zero at the 5 percent level, but not significantly different from one another. Both coefficients measure the extent of work in a year. The fact that they are not significantly different indicates that there appears to be no premium associated with full-time work, over and above the return to larger work effort.

Other specifications were tried but none substantially improved the quality of the estimates. Briefly, an education-experience interaction was included to capture the presumed positive correlation between formal school training and post-school on-the-job training. An age-experience interaction term was added to allow for the return to experience to depend upon age. In each case the t-statistic on these variables was not greater than one. Finally, the log of experience plus one was replaced by the arithmetic value of experience and its square. Both experience and its square were significant at the 5 percent level. Their respective coefficients were positive and negative, indicating a declining growth rate in wages from experience. The life-cycle wage-experience profile became negative at 25 years of experience. This suggests the possibility that at high rates of experience human capital depreciates faster with use than with nonuse. But it could also be due solely to the functional form which necessitates a downturn at some level of experience if growth rates in wages decline with experience. The R-squared in the two equations was virtually identical. Choosing between these two equations is a matter of judgment. We choose to rely on the logarithmic specification.

The interpretation of the parameters in our specification (equation 4.8) should be made clear. The coefficient on age, if cohort effects are ignored, measures the depreciation rate on wages resulting from spending a year out of the labor force. The coefficient on education measures the effect of an additional year of schooling relative to spending an additional year not at work. The effect of spending at least an additional year in the labor force for a woman of a given age is given by $\beta_5/EXP+1$. Growth rates in wages, therefore, are assumed to decline with additional years of experience. To compute the rise in wages resulting from additional years of experience as the woman moves through her life cycle, we must include the fact that she is aging. The life-cycle growth rate wages for a woman who continuously works full time is equal to $\beta_4 + \beta_5/EXP+1$.

The parameter estimates of the wage equation are reported in Table 9.

Table 9

WAGE EQUATION PARAMETER ESTIMATES

Variable	Coefficient	Standard Error
INTERCEPT	-.419	.133
SOUTH	-.051	.042
CITY	.104	.039
EDUCATION	.098	.009
AGE	-.009	.002
LEXP	.314	.133
R^2	.238	
F	63.58	

The estimated effect of education indicates that an additional year of schooling adds 10 percent to the woman's wage. This point estimate is slightly higher than estimates reported by Mincer and Polachek (1974), Heckman (1974), and Cogan (1977), all of whom used the National Longitudinal Survey of Mature Women. They estimate an effect of around 8 percent. It is also similar to the estimate reported by Smith (1978), who used Census data. The stability of this point estimate across specifications of the wage equation and samples is encouraging.

The estimated effect of a year of labor force experience on wages is 4 percent, when evaluated at the mean level of experience in the data. This estimate also conforms to estimates of earlier research. Heckman (1974) and Cogan (1977) estimate returns to experience of slightly over 2 percent. However, the measure of experience used in their work is the number of years in which the woman worked at least six months during the year. As a result, their estimates are interpreted as the effect of spending less than a full year in the labor force, whereas our estimate is interpreted as the effect of a year of full-time work. If women who work at least six months of a given year work, on average, two-thirds of a year, the estimates of Heckman and Cogan imply a 3 percent return to a year of full-time work. The estimated return to experience reported by Mincer and Polachek (1974) is much lower, approximately 1.2 percent.

The estimated effect of experience declines markedly with the level of experience. For instance, the effect of the fifth year of labor force experience adds 5.2 percent to the woman's wage, while the tenth year adds about half as much, 2.8 percent. This result is not due entirely to the functional form of the experience variable (e.g., the natural logarithm of experience). Essentially, the same decline in the impact of experience was obtained when experience and its square were entered as regressions.

If there were no differences in wage profiles among women of different birth cohorts, the coefficient on age would measure the rate of depreciation associated with time spent out of the labor force.

The coefficient on age is .9 percent. Thus, if a woman were to spend ten years out of the labor force, her wage at re-entry into the labor force would be 9 percent lower than it was at the time she exited. This estimate seems low, especially when one considers the probable impact of cohort effects. If wage profiles of younger cohorts are higher than those of older cohorts, as Lillard and Weiss (forthcoming) hypothesize, our estimated depreciation rate would overstate the true depreciation rate.

There are two plausible reasons why we estimate low depreciation rates. One is that younger cohorts of women have lower rather than higher wage profiles relative to older cohorts. This would bias the age coefficient in a positive direction and result in an understatement of the depreciation rate. There is some reason to believe that this is the case. Welch (1978) has hypothesized that the size of the cohort (negatively) affects the level of the wage profiles of members of the cohort. The youngest women in our data are members of a relatively large (baby boom) cohort. An alternative explanation is that employers, given a fixed cost of hiring, are likely to pay a wage premium for women who have completed their fertility and, as a consequence, are less likely to leave a given job. Such a wage premium may dominate the depreciation in wages resulting from time spent out of the labor force.

The coefficients on the residence dummy variables have their usual signs and magnitudes. Women who live in urban areas with populations over 100,000 earn 10.4 percent more than those who do not. Women living in the South earn 5.1 percent less than women living in other regions of the country, though this estimate is not statistically significant.

Figures 2 and 3 depict age-wage profiles of women with different amounts of accumulated work experience. In Figure 2, the profile labelled complete participation describes the wage profile of a woman with 12 years of education who works full time during every year of her post-school life cycle. Similarly, the half-time participation profile describes the wage profile of a woman with 12 years of

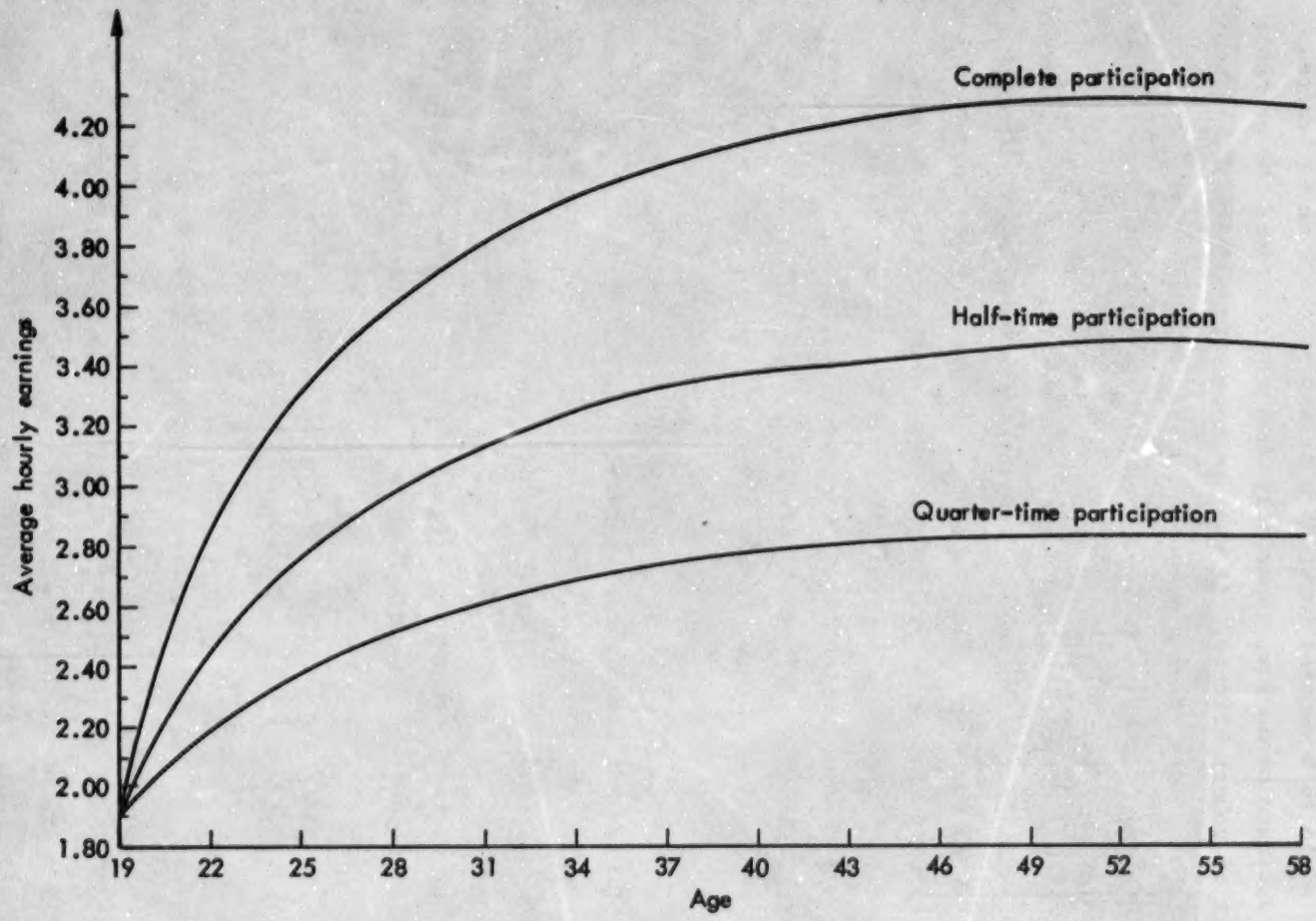

Fig. 2—Simulated age-wage profiles by proportion of year worked; white married women with 12 years of education

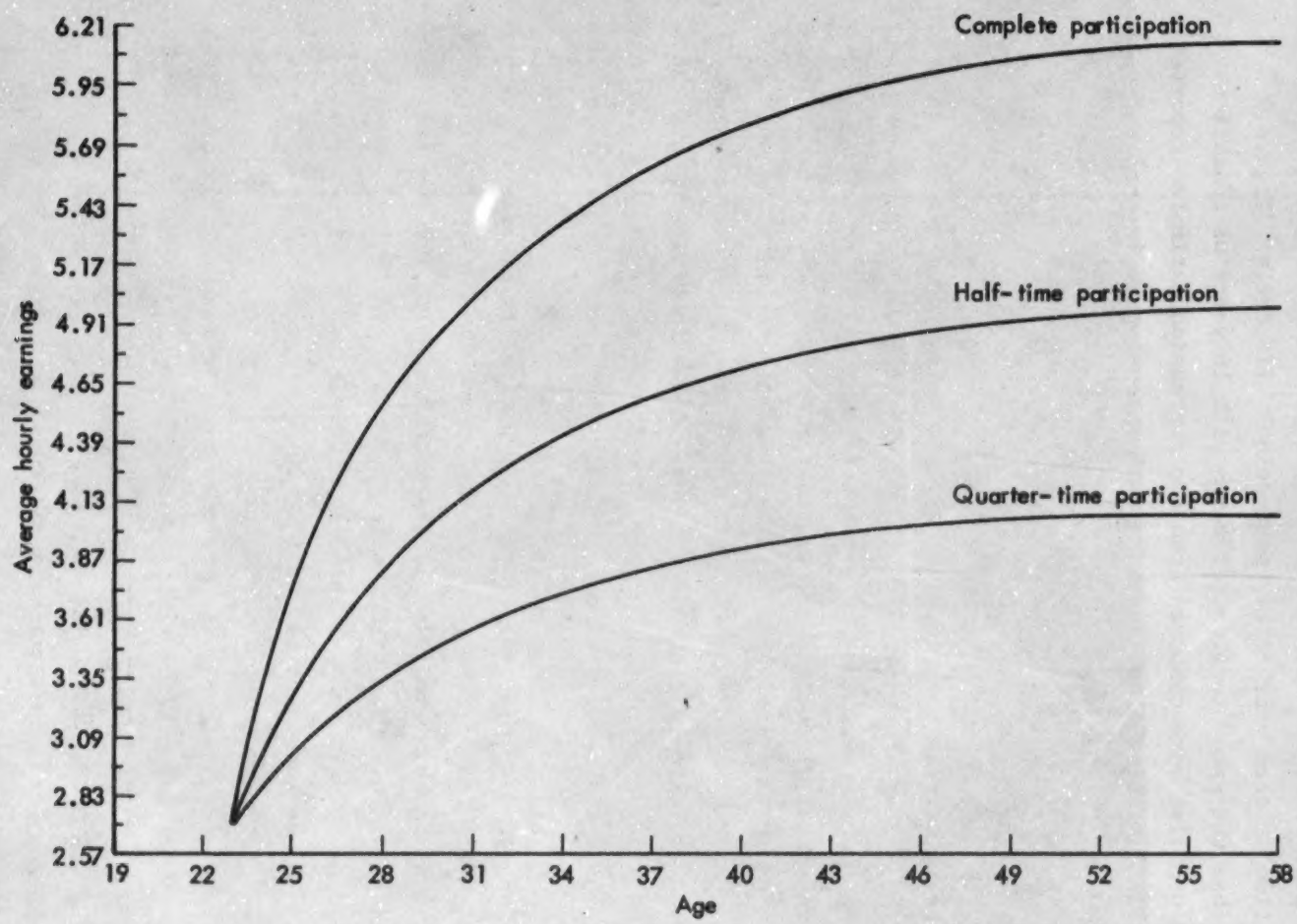

Fig. 3—Simulated age-wage profiles by proportion of year worked; white married women with 16 years of education

education who works full time during every year of her post-school life cycle. Similarly, the half-time participation profile describes the wage profile of a woman with 12 years of schooling who works only half-time in each year of her post-school life cycle. Figure 3 describes the same profiles for women with 16 years of education. These profiles were computed from the regression estimates reported in Table 8. Using these estimates, the growth rate in real wages is

$$\frac{dW}{dA} = F \frac{.314}{X+1} - .009\ W \tag{4.9}$$

where F is the fraction of each year the woman spends in the labor force, X is her accumulated years of full-time work experience, and W is her market wage rate at a point in time.

The total growth in real wages from age of entry into the labor force to age 55 among women with 12 years of education who work full time in each year is slightly over 150 percent. Among women with 12 years of education who spend each year in the labor force, but who work only one-half time, the total growth is 100 percent, and among women who spend only one-fourth of each year in the labor force, it is 62 percent.

The age-wage profiles reach their peak at later ages the larger the extent of labor market participation and the higher the level of education. Among women with 12 years of education who are full-time participants in the labor force in each year, the profile peaks at age 55; among women who are half-time participants, it peaks at age 54; and among women who are quarter-time participants, it peaks at age 52. For women with 16 years of education, the full-time work profile peaks at age 59, the one-half profile peaks at age 58, and the quarter-time profile peaks at age 56.

4.5 CHILDREN AND WAGE RATES OF WOMEN

By combining the estimates of the effect of children on the mother's labor supply reported in Section 3 with estimates of the impact of experience on wages reported in this section, we can obtain

estimates of the reduction in wages that results from child rearing. Figure 4 depicts age-wage profiles of two representative women--one who bears no children and the other who has the mean characteristics in our data. The average woman in the data begins her family at age 22, has three children during the course of her life, and spaces these children three years apart.[1] Wages of the woman who bears no children are higher and initially rise at a faster rate than the child-bearing woman. The woman with children actually suffers a decline in her wage rate from age 28 to age 31. The largest difference in wage rates occurs at age 32, when the wage of the woman with no children is 22 percent higher than the wage of the woman with children. The average wage over the life cycle of the woman with no children is 13 percent higher than the average wage of the woman with children.

These results for the "average" woman are actually due to diverse effects of numbers, timing, and spacing of children on wage rates. Figure 5 depicts the relationship between wages and spacing of births. Shown are age-wage profiles for women who have three children over their lifetime, spaced, alternatively, 2, 4, and 6 years apart (beginning at age 19). Women who space their children closer together have lower wage rates during the child-bearing years as their accumulated experience lags behind that of women who are spacing their children further apart. After fertility is completed, wage growth accelerates, especially for women who space their children close together. This is because the rate of growth of wages is negatively related to the level of accumulated experience and because women who space their children close together have lower levels of accumulated experience at the time of completed fertility. The maximum differential between women who space their children two years apart and those who space their children six years apart is 17 percent and occurs at age 25.

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[1] Other characteristics of the two women are equal.

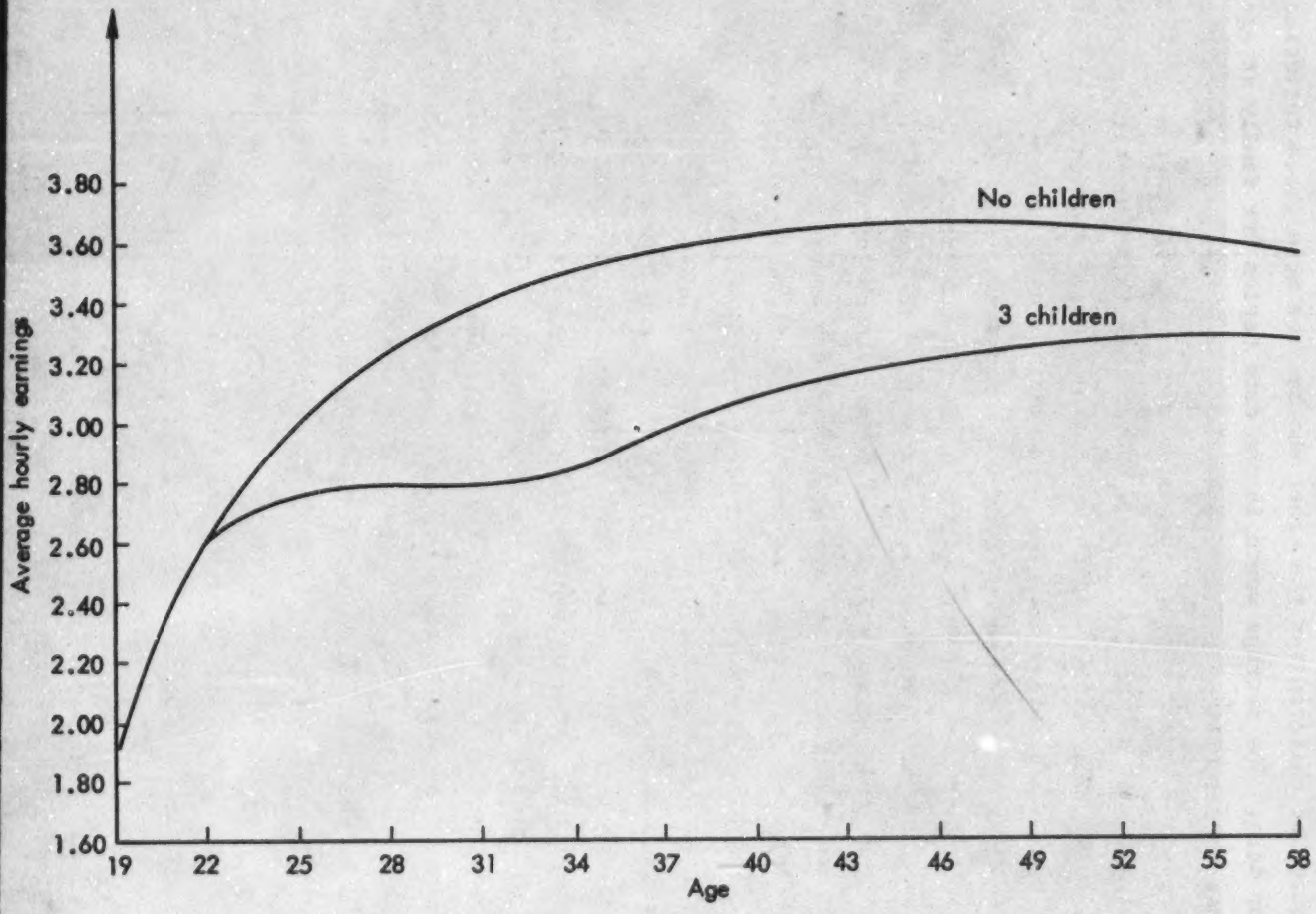

Fig. 4—Simulated effects of children on age-wage profiles

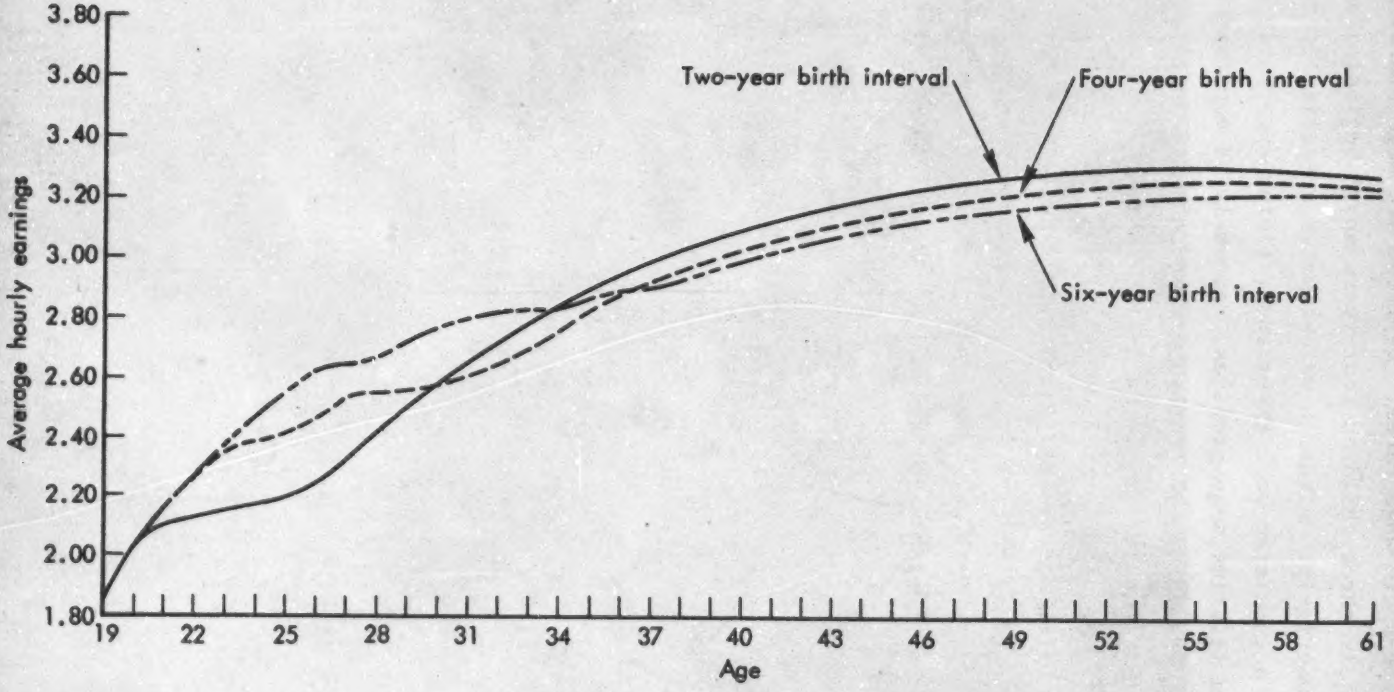

Fig. 5—Spacing and life-cycle wages

Figure 6 represents the relationship between timing of the first birth and life-cycle wage rates. Depicted are wage profiles of women who space their three children four years apart beginning at ages 19 and 22, respectively. Although we estimated a negligible timing effect on lifetime accumulated work experience, timing apparently does have an effect on lifetime wage profiles. Accumulated experience is significantly lower during the early years of the life cycle for women whose first birth occurs early in life. Therefore, during these years their market wage offers are lower as well. In later years, women whose first birth occurs earlier in life tend to work more; as accumulated experience of the two groups converge, so do wage rates.

Figure 7 presents age-wage profiles for women who have 0, 1, 2, and 3 children, beginning at age 19, and spaced (where relevant) four years apart.

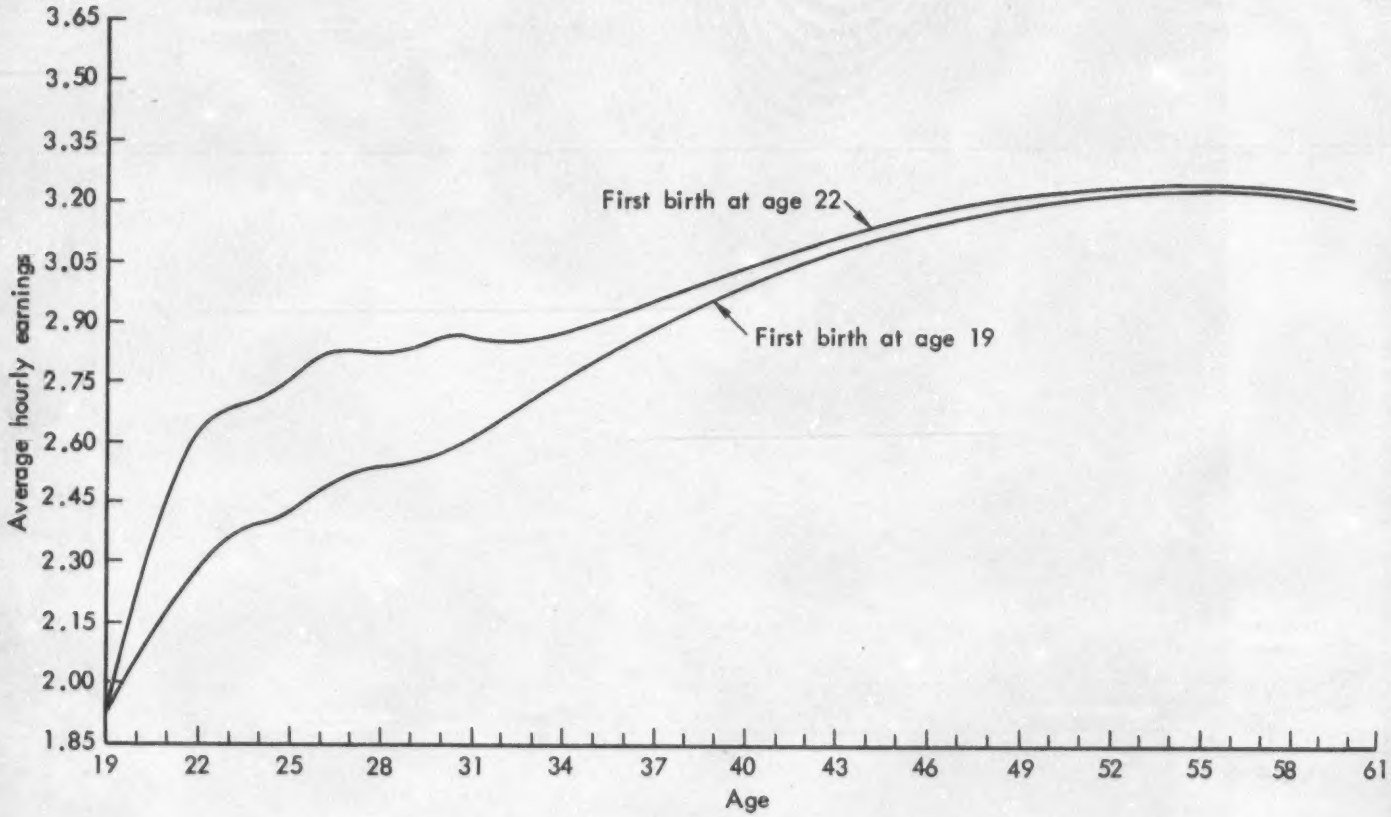

Fig. 6—Timing and life-cycle wages

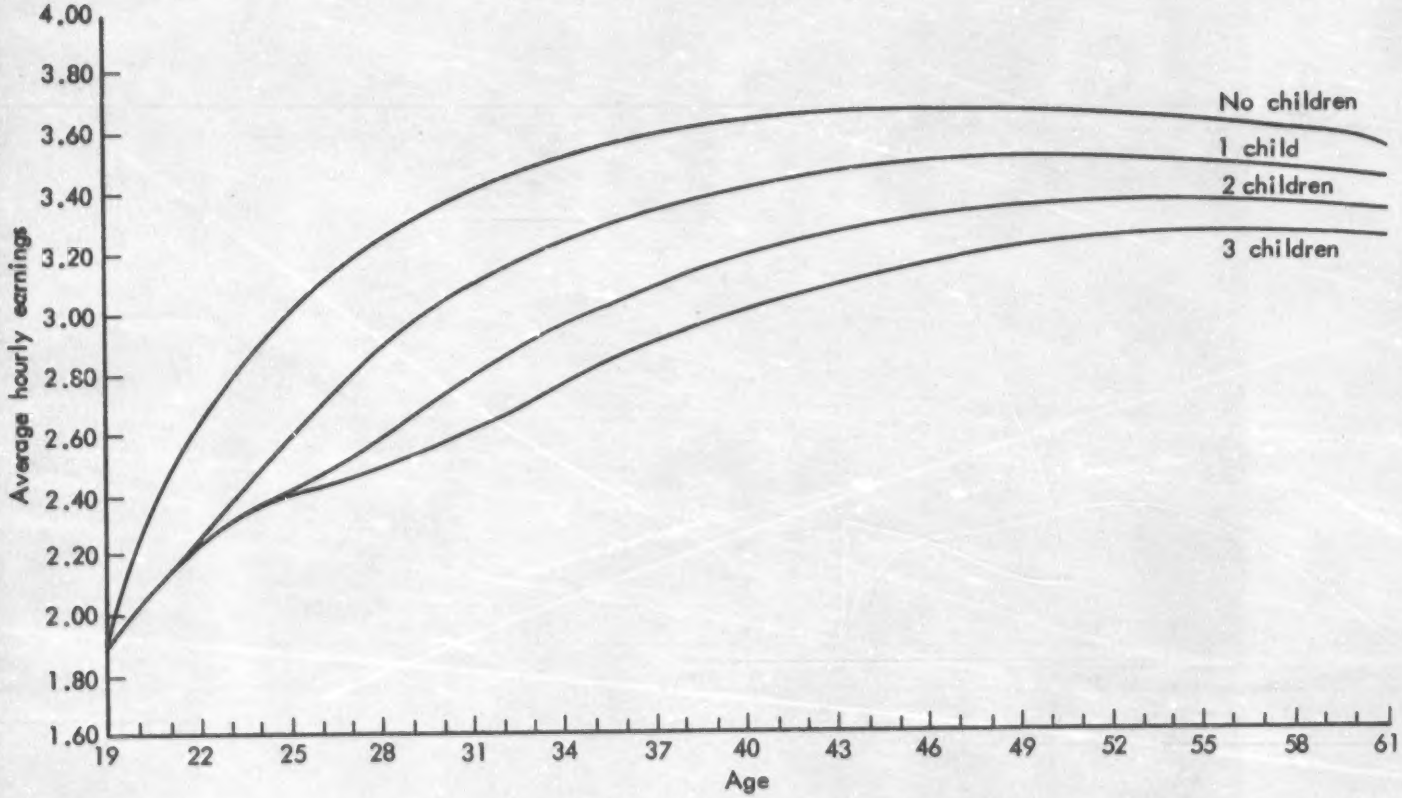

Fig. 7—Numbers of children and life-cycle wages

5. CONCLUSIONS

In this report we have attempted to relate the number, timing, and spacing of children to accumulated work experience and wage rates of women. Our principal findings are as follows:

- The number of children a woman rears has an important deterrent effect on her accumulated work experience. A woman who has one child will work, on average, about 2.5 years less over her lifetime than a woman with no children. A second child results in an additional two years withdrawn from the labor market, and a third one additional year.
- Larger birth intervals (holding constant number and timing of children) are associated with larger reductions in lifetime work effort. The magnitude of this effect appears to increase with the number of children. For a woman with three children who begins her family at age 19, we estimate that an increase in the average birth interval from two years to four years reduces lifetime labor market work by about one year. An increase in the average birth interval from four years to six years reduces lifetime labor market work by an additional one-half year.
- The timing of the first birth (the age at which a woman has her first child) does not appreciably affect lifetime labor market experience.
- Accumulated work experience is a strong determinant of market wage rates. Evaluated at the mean experience level in the data (7 years), the rate of return to an additional year of experience was estimated to be about 4 percent.

- Child rearing, through its effect on accumulated work experience, has a substantial effect on wage rates. The average wage rate over the life cycle of a woman with no children is 13 percent higher than that of a woman who has three children, spaced three years apart, beginning at age 22 (the "average" woman in our sample). The maximum difference in wage rates is 22 percent and occurs at age 32.

These findings must be considered tentative because of certain restrictive assumptions that we found necessary to make. These assumptions and their implications are summarized as follows:

- Child rearing is held to be exogenous with respect to both the labor supply decision and the market wage offer. This assumption is necessary to order to identify the parameters of the labor supply and market wage equations. To the extent that exogeneity fails to hold in the data, our estimates are biased.
- We have assumed that intertemporal variations in hours of work are not state dependent. To the extent that the amount of time a woman works in a given year is a determinant of subsequent labor supply decisions, then we have underestimated the effects of children on accumulated work experience. State dependence can be incorporated into the analysis only by using costly maximum likelihood estimators.
- In order to interpret our estimated age-related parameters as reflecting life-cycle behavioral relationships, it was necessary to assume that labor supply and wage equation parameters do not differ across women of different ages. In order to distinguish such cohort and true life-cycle effects it would be necessary (and interesting) to make use of the

longitudinal nature of the Income Dynamics Panel data. Such an investigation should be of high priority in future work done in this area.

Appendix A

TWO-STAGE LEAST SQUARES ESTIMATES OF WAGE EQUATION

Variable	Coefficient	Standard Error
Intercept	-.404	.141
South	-.052	.042
City	.103	.039
Education	.097	.009
Age	-.010	.005
Lexp	.339	.087
F	63.41	

Appendix B

WAGE EQUATION PARAMETER ESTIMATES CORRECTED FOR SELECTIVITY BIAS[1]

Variable	Coefficient	Standard Error
Intercept	-.131	.183
South	-.061	.042
City	.109	.039
Education	.081	.011
Age	-.007	.002
Lexp	.315	.031
$\frac{f(I/\sigma_u)}{p(I/\sigma_u)}$	-.232	.100

[1]The standard errors should be regarded with caution since no account in estimation is taken of heteroscedasticity.

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